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THE EFFECT OF INCENTIVE-INDUCED RETIREMENT ON SPOUSAL RETIREMENT RATES: EVIDENCE FROM A NATURAL EXPERIMENT

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We identify, quantify, and explain the impact of incentive-induced early retirement (ER) of husbands on their wives' probability to retire within 1 year, using administrative data from the Netherlands. Our identification strategy is based on a policy intervention by which targeted individuals working at the central government level became unexpectedly and temporarily eligible for very generous ER benefits. This retirement window of opportunity implied for interested workers that transitions from the current job into full retirement had to be effected swiftly and irreversibly. We find that induced ER of husbands increased their wives' probability to retire by 10 percentage points. This is a strong and robust local average treatment effect. Partly, the effect runs through wives at ages when they may have been eligible for ER programs themselves. (*JEL* C26, J26, J120, J140)

I. INTRODUCTION

Large changes across Organisation for Economic Co-operation and Development countries are being observed in labor force participation and retirement patterns of older workers. In particular, average ages at final labor force withdrawal are higher now than they used to be in previous decades, although strong inter-country differences remain (Blundell, Bozio, and Laroque 2013; Schirle 2008). These changing patterns have partly been ascribed to changes in institutions such as restricted access to early retirement (ER), or disability insurance (DI). Understanding the way individuals make labor supply and retirement decisions is crucial for designing effective policies that are meant to change behavior. Traditional microeconomic retirement models that can guide policy makers in their choice typically focus on individual worker decisions in isolation and study the decision process as a function of age, income, health, wealth, and financial or tax incentives (Berkovec and Stern 1991; Gustman and Steinmeier 1986; Lumsdaine, Stock, and Wise 1992; Rust and Phelan 1997). As a recent strand of literature has emphasized, however, labor force decisions of spouses are interrelated through various channels. Ignoring effects on labor supply of policy-targeted individuals' spouses may have direct implications for the evaluation of such policy measures (Blau and Gilleskie 2006; Gustman and Steinmeier 2009; Van Der Klaauw and Wolpin 2008).

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ABBREVIATIONS

ELSA: English Longitudinal Study of Aging
HRS: Health And Retirement Study
IV: Instrumental Variable
LATE: Local Average Treatment Effect
OLS: Ordinary Least Squares

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This paper studies whether shocks to ER incentives, relevant for certain employees in the age group 55–60, actually led to earlier labor force outflow, and whether it may have had spillover effects on retirement behavior of the spouses of those targeted. The policy change that we exploit became effective in 2005 for certain birth cohorts of civil servants employed for more than 10 years by the Dutch central government. These individuals were offered the opportunity to retire during the year 2005, by a temporary reduction of the ER eligibility age. For our empirical work, we focus on dual-earner couples in which the husband did not and the wife did work in the private sector, such as to rule out coincidental treatment of wives through the same reform.

After studying the intention-to-treat effect by means of a difference-in-difference specification, we employ an instrumental variable (IV) approach to estimate the effect of induced retirement on spousal retirement status. The changed ER incentive for civil servants serves as an exogenous shifter of retirement. In other words, our identification strategy relies on an unforeseen window of opportunity that opened up for a specific group of workers, and not on fixed and well-known age rules. Identification makes use of private sector workers that were unaffected by the same reform.

A worker's retirement can affect spousal retirement through various channels and in different directions. Upon the retirement of an individual, the spouse may postpone her or his own retirement to prevent drops in household income. Alternatively, spouses may retire simultaneously, because they enjoy spending leisure time together such that retirement of one spouse induces the retirement of the other. Without instrumenting retirement status, estimates of the effect of retirement on spousal retirement status are likely to be biased. Bias may arise if men and women with similar preferences tend to match, or if spouses with a preference for joint retirement select themselves into jobs that enable them to do so.

The wife's probability to retire is the main dependent variable in our IV model. We employ the mentioned policy change as an instrument for husband's retirement status, providing variation in husbands' retirement rates across public sector (treated) and private sector (control) employees, and over time. Lack of a sufficient number of observations on wives that were induced to retire through the reform and who had working husbands precludes us from providing a similar analysis for women's retirement and possible spillover effects on husbands.[1]

Using Dutch administrative micro data at the population level, we find reliable evidence for an effect of induced ER for husbands in the first stage. A reduced form difference-in-difference analysis of the husband's changed financial incentives to retire early shows clear and nearly immediate spillover effects on his wife's retirement decision. The second stage of our IV approach confirms the robust and positive impact of the husband's predicted retirement status on his wife's retirement.

According to our central estimates, retirement of male civil servants induced by the policy change led to a jump in the probability of their wives to retire within 12 months by 10 percentage points. An alternative reduced form analysis of the wives' ages shows that the effect is driven by husbands whose wives may have been eligible for regular ER benefits themselves. This result may be explained by the presence of leisure complementarities. If husbands are induced to retire, wives who do not forego a financially attractive retirement opportunity may quit working as well, so as to spend their leisure time together with their spouses. Husbands with wives who reach the eligibility age for regular ER benefits may also be more likely to accept an ER offer themselves.

We directly contribute to a recent empirical literature that estimates the effect of incentive-induced retirement of one partner on retirement status of the spouse. The identification methods used in this literature rely on the relevant incentives being exogenous sources of variation in the retirement rates. Banks, Blundell, and Casanova Rivas (2010) is the study most comparable to ours, based on survey data, however. They find that spousal decisions can be influenced by retirement shifts of their partners. They use variation in eligibility ages for ER benefits between countries as source of variation in the probability to retire. We exploit variation in retirement rates across sector, age, and time. Whereas it

1. One underlying cause is the relatively weak labor market attachment of women, in particular for the ages and birth year cohorts affected by the policy change we study. Husbands having reached eligibility ages for ER benefits is a restricting factor as well. Husbands of wives in the age category 55–60 were typically in their late 50s or early 60s. Early retirement arrangements in the private sector typically offered early retirement as of age 60, so that many of those husbands in their early 60s were already retired.

is common to rely on fixed age rules in related papers, a remaining objection is that workers may retire or postpone retirement while anticipating the retirement of their spouses. Such anticipation effects could bias the estimated treatment effect upwards or downwards. Our research design, relying on a shock to eligibility conditions, avoids this in principle.

Our first contribution is to use strong instruments that provide exogenous (unanticipated) variation in retirement rates of husbands. Second, we use administrative data that include end dates of jobs and that allow us to observe the precise within-couple sequencing of retirement. This is critical in order to rule out that our estimates are influenced by behavior of wives that actually retired earlier than their husbands. The latter effect cannot necessarily be ruled out in studies that are based on biennial survey data, posing a potential threat to identification. Third, as we have access to data covering the entire population, we can focus on individuals from a very narrow age range with a specific career.

The rest of the paper is set up as follows. Section II reviews the relevant related literature and Section III describes the institutional environment, including the policy change that we exploit. Section IV explains and describes the data. Section V delineates the identification strategy we use for estimating the causal effect of husband's retirement on wife's retirement status, and discusses results. Section VI concludes.

II. LITERATURE REVIEW

Structural models of joint retirement assume that husbands and wives make separate retirement decisions and have their own preferences. Papers such as Blau and Gilleskie (2006) and Van Der Klaauw and Wolpin (2008) carefully model the incentives provided by social security rules, and specify stochastic processes for wages, health, and survival. The main effect of husbands' on wives' retirement choices runs through the household budget constraint in those papers. The household budget constraint is not the only channel inducing dependence, however. Spouses enjoy spending time together, that is, spousal preferences directly depend on one another (Gustman and Steinmeier 2000, 2009; Hurd 1990). Gustman and Steinmeier (2000), studying retirement choices assuming absence of uncertainty, find that spouses coordinate retirement decisions and that this coordination is motivated by leisure complementarities rather than financial incentives provided by the household budget constraint. Casanova Rivas (2010) does take into account uncertainty regarding future income, health, and survival. In line with Gustman and Steinmeier (2000), she finds evidence for leisure complementarities. Michaud and Vermeulen (2011) employ a structural approach incorporating intrahousehold interactions and find evidence for leisure complementarities as well.

A second strand of empirical literature estimates the effect of incentive-induced retirement of one partner on retirement status of the spouse. Zweimueller, Winter-Ebmer, and Falkinger (1996) estimate a bivariate probit model, where the dependent variables are the retirement statuses of the husband and the wife. The retirement status of each of the partners depends on the social security variables of both spouses. Using data from the Austrian Microzensus, the authors find that husbands responded to a change in the minimum retirement age of their wives whereas wives did not respond to a change in the minimum retirement age of their husbands. Baker (2002) estimates the effect of a decrease in the eligibility age for age-related income security benefits for workers on labor force participation of spouses in Canada, for a sample of workers who were younger than their spouses. He uses data from the Canadian Survey of Consumer Finances. The author finds that eligibility for the age-related benefits for wives was associated with a 6–7 percentage point decrease in labor force participation of husbands. He does not find an association between eligibility for the age-related benefits of husbands and labor force participation of wives.

Banks, Blundell, and Casanova Rivas (2010) employ a difference-in-difference approach, exploiting the difference in eligibility ages for ER benefits between England and the United States as a source of variation in retirement status. They use data from 2002 to 2004 waves of the Health and Retirement Study (HRS) for the United States and the English Longitudinal Study of Aging (ELSA) for England. The authors find that men in England were 14–20 percentage points more likely to retire within 2 years when their wives reached the ER age than comparable men in the United States. These effects are found for three subsamples of men who were at least as old, at least 1 year older, or at least 2 years older than their wives. The authors do not find an effect if they do not restrict their sample based on age differences between men and their wives.

Stancanelli (2017) employs a sharp Regression Discontinuity design, using the youngest ER age in France (the age of 60) as the discontinuity in the probability to retire. She uses data from the French Labor Force Surveys (LFS), and finds a negative effect of retirement on the partner's hours worked, both for males and females. Lalive and Parrotta (2017) employ a double regression discontinuity design on Swiss data and find that wives reduce their labor force participation by 2–3 percentage points when their husbands reach pension eligibility. They do not find such spillovers for husbands.

Selin (2017) has a similar finding for spillovers for husbands, exploiting a pension reform in Sweden. Broad categories of local government workers were eligible for full pension benefits at the age of 63 until 2001, but lost eligibility in the 2001 reform. The author estimates a difference-in-difference model using administrative data and does not find evidence for a response in husbands' retirement behavior to their wives' changed retirement incentives.

III. INSTITUTIONAL BACKGROUND AND POLICY CHANGE

We shall focus on targeted incentives to retire early that became available to a group of civil servants in the Netherlands.[2] The Dutch retirement system foresees retirement at the standard age (for both men and women) of 65. Actual average ages of entering retirement have been considerably lower, however, because of the widespread use of ER arrangements in virtually all sectors of the economy.[3]

The Dutch pension system rests on three pillars. The first pillar is the public old-age pension system, which is financed on a pay-as-you-go basis. The second pillar consists of occupational pensions. Regular pensions are funded. ER pensions are financed on a pay-as-you-go basis. The third pillar consists of private provisions. We study the period around 2005. At that time, most occupational pension funds offered ER arrangements, allowing for retirement below the standard age of 65. Individual pension funds set their own entry age for ER benefit receipt (on average just above the age of 60) and stipulated eligibility criteria. For instance, the public sector pension fund offered ER arrangements as of the ages 61 or 62 onwards, that allowed continuing into regular pension benefit receipt from the age of 65 on. Retirement before reaching the ER eligibility age was financially very unattractive. The ER benefits were financed with premiums, separately from the regular pension benefits, and because of a favorable tax treatment claiming ER benefits upon eligibility was actuarially highly beneficial. We use a temporary decrease in the ER eligibility age for civil servants as a source of exogenous variation to estimate the impact of incentive-induced ER of husbands on their wives' probability to retire within 1 year.[4]

This temporary decrease—which we refer to as "the ER window"—was announced in April 2004.[5] The government aimed at reducing and restructuring public sector tasks. Ministries (departments) and their agencies qualified to offer certain civil servants additional possibilities for ER during the first 11 months of 2005, under the provision that forced layoff (in the wake of reorganizations) of younger civil servants could be avoided.[6] In practice, this was implemented by offering ER collectively to all or none of the qualifying workers within a qualifying department (Dutch Government 2004). This aspect is vital for the internal validity of our identification strategy, because if employers had offered the incentive selectively to such husbands who wished to spend time with their retiring wives, for example, the opening of the ER window would be endogenous to the retirement status of wives. The ER window offered gross retirement benefits that could be up to 70% of workers' average pay (mid-career salary), corresponding to benefit levels in other ER programs.

Civil servants faced several eligibility criteria for leaving within the ER window (Dutch Government 2004, 2005), in particular age and career-related norms. First, they had to be at least 55 years old when retiring. Second, they had to have been continuously employed as a civil servant during the 10 years prior to retiring. They were also required to have continuously

2. See also Bloemen, Hochguertel, and Zweerink (2017).

3. A concise description of the Dutch pension system and existing early retirement arrangements for civil servants valid at the time of our data can be found in Appendix A: There, we also consider the role of competing pathways out of the labor force, in particular DI.

4. We consider a 12 month period, not a calendar year.

5. The term "retirement window" has been used before in work studying similar types of policy reforms, for example, Lumsdaine, Stock, and Wise (1992), Börsch-Supan and Schnabel (1998), Lumsdaine and Mitchell (1999) and Coe et al. (2012).

6. Ministries and their agencies typically have thousands of employees.

contributed to the public sector pension fund during the 10 years prior to ER. Importantly, these requirements prevent self-selection into the public sector of workers who would like to retire early.

Employers then had to decide before January 1, 2005 whether or not to open the ER window as of January 1, 2005, and eligible civil servants were not allowed to retire later than December 1, 2005. These stipulations hence meant that retirement decisions had to be taken swiftly and irreversibly.

As the regular public sector ER arrangement continued to be offered from ages 61 or 62 on, the relevant treatment group for our natural experiment thus contains individuals aged 55–60. We compare with private sector workers of the same age (control group), where such a window of opportunity did not arise, and we compare before and during the treatment year 2005.

It is important to realize, however, that not all age groups within our treatment group will have been exposed to an equally strong incentive to retire. There are two reasons. One is a maximum duration of ER benefits of 8 years under the window. There was no further provision for a possible gap between expiration of ER benefits and start of receipt of regular pension benefits from age 65 on. Civil servants aged 57 or older at the onset of ER were not facing a coverage gap; those aged 56 or younger would see their ER benefits exhausted 1 or 2 years prior to normal pension receipt eligibility. The other reason is differential accrual possibilities for regular pension benefits. In particular, civil servants born before January 1, 1948 could continue accruing pension claims at a rate of 50% at the expense of the employer for a maximum of 4 years. Civil servants born on or later than January 1, 1948, that is, civil servants who were in the age category 55–57 in 2005, did not have this opportunity. In summary, the ER window was thus very attractive for civil servants aged 58 and older, somewhat less attractive for those aged 57, but substantially less attractive for civil servants aged 55 or 56. In our main analysis we therefore differentiate our instrument by age, allowing for heterogeneous treatment effects.

IV. DATA

We use administrative data collected and prepared for research purposes by Statistics Netherlands. The observations we use cover the period 2000–2005 and include variables on job and personal characteristics.[7] The job characteristics data provide information on all jobs that any registered individual has held. The job information includes job spells (precise start and end dates per job), the industry code, and the annual wage. The personal characteristics data cover the entire population, allow linking partners in a couple, and contain information on demographic variables such as nationality, marital status, birth year, and birth month.

Important for our identification strategy is a classification of workers into treatment and control groups. We base those on industry affiliation observed in the data. The ER window was opened for civil servants employed by qualifying central government organizations. We do not observe ER window offers at individual worker level, nor do we directly observe at which department a civil servant worked. This implies that we cannot observe whether a civil servant who did not retire rejected the ER offer or simply was not offered the ER window. Hence the "treatment" group as defined in our data is larger than the "true" treatment group.[8]

We define a baseline sample for the years $t = 2000, \ldots, 2005$. The sample selection is guided by the qualification criteria for the reform. The sample consists of opposite-sex couples that were married during the entire year t.[9] Further, husbands needed to be in the treatment-relevant age range of 55–60 (measured on December 31st of year t), and have been continuously employed for the 10 years prior to January 1st of year t. The treatment group then has husbands that were central government civil servants, the control or comparison group has husbands that were private sector workers. In all couples of our sample, the wife was employed in the private sector on January 1st of year t. Our selection ensures that husbands in the treatment group could have been eligible for

7. The original file names are *Doodsoorzaken (2000–2005)*, *Landelijke Medische Registratie* (*LMR, 1999–2004*), *SSB Banen (1999–2008)*, *SSB Personen (2000–2005)* and *PARTNERBUS (2010)*. We use vital statistics to establish that a nonworking individual is not deceased. Unfortunately, for the years we are interested in, there are no data available on, for example, taxable financial wealth. We do not use observations for years after 2005 because there were some key variables such as wage income that were measured differently after 2005.

8. We do not know how much smaller the "true" treatment group is than the observed treatment group.

9. Being married includes having a registered partnership. Registered partnership refers to partnerships enjoying legal status similar to marriage. Being married excludes cohabitation without being married or without having a registered partnership.

entering the ER window for civil servants, and wives could not have been. As additional criterion we exclude observations on workers who earned less than 20,000 euros in the year prior to year t, so as to ensure that workers had a strong labor force attachment.[10] Individuals with a weaker labor force attachment may be found out of employment, even if no formal retirement choice is involved. We subject the data selection criteria to a battery of sensitivity checks in Appendix B1.

We define retirement as leaving a job and not having started working again before January 1, 2009. We use retirement of the first retiring member of the couple as an absorbing state. This implies that we do not use observations for years after a member of the couple had retired. By doing so, we ensure that all observations on retirement in our sample concern transitions into retirement rather than retirement spells that started in previous years. We use about 100,000 observations for our analysis.

Table 1 shows descriptive statistics for our baseline sample, split by treatment and control group and by treatment year (Panels a and b) and pretreatment period (Panels c and d). The income variables measure the total wage income the wife and her husband earned (in year $t-1$, and expressed in tens of thousands of deflated euros). Wage income is a relevant factor for retirement decisions as it determines the level of pension benefits. It is also correlated with the number of hours worked. To capture health effects on retirement with our administrative records, we include hospitalization status information. Specifically, we use dummy variables indicating whether the wife or husband had been hospitalized during the previous calendar year ($t-1$). For instance, we may pick up retirement of an individual that cares for a sick spouse. Further, we include dummy variables indicating whether the wife or husband had the Dutch nationality. Nationality can pick up immigration status and be correlated with the number of contribution years at pension funds and the social security system, and so with the overall level of retirement benefits. There can also be cultural reasons why people with different nationalities may make different retirement decisions.

Table 1 shows that across periods (pretreatment and treatment) husbands and wives are on average very similar in terms of age, wage income, hospitalization status, and having the Dutch nationality. This is the case for both husbands employed as civil servants and husbands employed in the private sector. Our results of t test indicate significant differences in means for some characteristics, but these differences are typically small in magnitude. For slightly more than 20% of all our sample couples, the husband was employed as a civil servant, which is a sizeable group. Conversely, about 18% of husbands married to a private-sector worker were civil servants.

The couples in our analysis are fairly representative for an average Dutch couple of the same age group. Dual working couples are not uncommon in the Netherlands; 78% of the men in the age category 55–60 in the Netherlands in 2005 were married. Of them 85% were employed, and 44% of those had employed spouses. Nine percent of these dual-earner couples consisted of a husband working as a civil servant and a wife being employed in the private sector. Civil servants represent a heterogeneous group of workers, making findings based on observations for civil servants also relevant for workers employed in many other industries.

V. EMPIRICAL APPROACHES AND RESULTS

A. *Introduction*

We use a difference-in-differences approach to first get an idea how the changed ER incentives may have impacted the labor force status of the male workers we consider. We provide graphical evidence and discuss the parallel trend assumption. To account for the fact that males in the upper age range of our age bracket, the 58–60 year olds, in fact receive a stronger incentive to retire early than the ones in the lower age range (55–57), we split the treatment group in two, allowing each to have their own average treatment effect. We then move to analyze how the wives of males in the two treatment groups behaved. We apply a difference-in-differences approach with the wife's retirement as an outcome, measuring the effect that is often referred to as intention-to-treat effect (ITTE).[11] Results, discussed in Section V.C, give reason to believe that wives' retirement behavior in particular is causally changed by the changed retirement behavior of their spouses (husbands). Accordingly, we are interested in measuring

10. As the annual minimum wage for full-time employment was 16,400 euros in 2005, our threshold is slightly higher than the minimum wage for full-time employment.

11. This approach is taken by Selin (2017) for the effect of female retirement on their husbands' participation.

TABLE 1
Descriptive Statistics (Baseline Sample)

	Panel a Control Group Husbands Employed in the Private Sector Year 2005		Panel b Treatment Group Husbands Employed as Civil Servants Year 2005		Difference-in-Means		
Variable	**Mean**	**Std. Dev.**	**Mean**	**Std. Dev.**	**Mean**	**Std. Err.**	***p* Value**
Wife's age	53.16	4.53	53.70	4.01	−0.541	0.078	0.000
Husband's age	57.08	1.59	57.08	1.59	0.003	0.028	0.922
Wife's wage income [$t-1$] (in 10,000s of Euros)	3.25	1.46	3.21	1.49	0.046	0.026	0.072
Husband's wage income [$t-1$] (in 10,000s of Euros)	4.80	2.52	4.64	1.57	0.154	0.042	0.000
Wife hospitalized [$t-1$]	0.047	0.211	0.048	0.215	−0.002	0.004	0.677
Husband hospitalized [$t-1$]	0.086	0.281	0.084	0.278	0.002	0.005	0.696
Wife Dutch nationality	0.871	0.336	0.890	0.313	−0.019	0.006	0.001
Husband Dutch nationality	0.864	0.342	0.903	0.296	−0.039	0.006	0.000
Wife's retirement	0.022	0.148	0.030	0.172	−0.008	0.003	0.002
Husband's retirement	0.059	0.235	0.119	0.324	−0.060	0.004	0.000
N	17,880		4,012				

	Panel c Husbands Employed in the Private Sector Years 2000–2004		Panel d Husbands Employed as Civil Servants Years 2000–2004		Difference-in-Means		
Variable	**Mean**	**Std. Dev.**	**Mean**	**Std. Dev.**	**Mean**	**Std. Err.**	***p* Value**
Wife's age	52.74	4.64	53.39	4.07	−0.644	0.044	0.000
Husband's age	56.87	1.59	56.87	1.59	0.000	0.016	1.000
Wife's wage income [$t-1$] (in 10,000s of Euros)	3.24	1.38	3.22	1.27	0.028	0.013	0.032
Husband's wage income [$t-1$] (in 10,000s of Euros)	4.86	2.49	4.66	1.52	0.200	0.023	0.000
Wife hospitalized [$t-1$]	0.041	0.197	0.045	0.206	−0.004	0.002	0.045
Husband hospitalized [$t-1$]	0.073	0.260	0.070	0.254	0.003	0.003	0.210
Wife Dutch nationality	0.862	0.345	0.889	0.314	−0.027	0.003	0.000
Husband Dutch nationality	0.865	0.342	0.896	0.306	−0.031	0.003	0.000
Wife's retirement	0.025	0.157	0.024	0.152	0.002	0.002	0.271
Husband's retirement	0.059	0.236	0.039	0.192	0.021	0.002	0.000
N	60,583		12,666				

Notes: This table shows descriptive statistics for husband-wife couples in which the wife was employed in the private sector, and in which the husband was aged 55–60 and was either employed as civil servant (public sector; treatment group; panels a and c) or employed in the private sector (control group; panels b and d).Treatment year is 2005 (panels a and b).

the local average treatment effect (LATE) of the impact of incentive-induced ER of the husband on the wife's probability to retire within 1 year. To measure this effect, we employ an IV approach (Section V.C). We instrument the retirement choice of the husband by the husband's eligibility for the ER window.

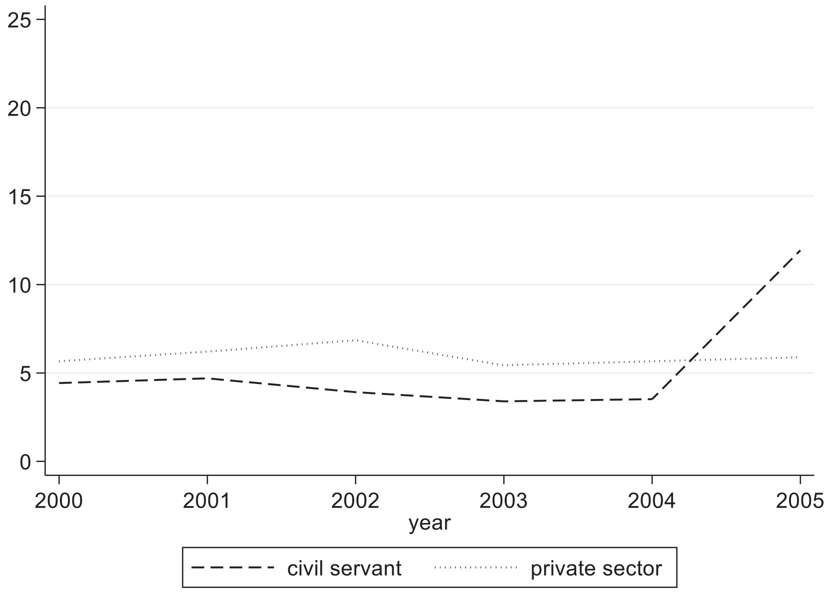
B. Difference-in-Difference Analysis

The difference-in-difference analysis will relate the policy change affecting husbands directly to the retirement status of the husband.

There are two assumptions that need to hold for the difference-in-difference approach to be valid: (1) treatment and control groups need to have experienced parallel trends in absence of treatment, and (2) there are no strong compositional changes between before and after reform measurements that affect groups differentially.

Regarding the second assumption, we found no sizeable differences in observables before and after the reform for the control and treatment group in Section IV. This finding gives us reason to believe that there were no strong compositional changes in unobservables that may bias our results later on. Also, the eligibility criteria

FIGURE 1
Husbands' Retirement Rates by Year and Sector (Percentages)



Notes: This figure shows average 12-months-ahead retirement rates for husbands, by husbands' sector of employment. Based on the baseline sample (Table 1).

for the ER window, in particular the criterion on continuous employment in the public sector during the prior 10 years, made sure that the reform did not trigger selection into the treatment group.

The first assumption can be checked graphically and tested formally. Figure 1 shows retirement rates for husbands in the age category 55–60 across years. There is a clear upwards jump of about 8 percentage points visible in 2005 for the treatment group, although there is no differential behavior between the control and treatment group during the pretreatment years. We employ a basic difference-in-difference model without covariates to formally verify the parallel trend assumption:

$$
\begin{aligned}
(1)\qquad H_{it} = \alpha_H &+ \sum_{j=2000}^{j=2004} \beta_{jH} \iota(t=j) \\
&+ \sum_{j=2000}^{j=2003} \gamma_{jH} \iota(t=j)\, C_{itH} \\
&+ \gamma_{2005,H} \iota(t=2005)\, C_{itH} \\
&+ \zeta_H C_{itH} + v_{it},
\end{aligned}
$$

where H_{it} is a dummy variable that is 1 if the husband of couple i retired in year t, H_{it} is 0 otherwise,[12] $\iota(.)$ is the indicator function, C_{itH} is a dummy variable indicating whether the husband was employed as a civil servant in year t, and 0 otherwise (employed in the private sector).[13] Pretreatment year 2004 serves as the baseline year for the interactions between the dummies for year and husband being employed as a civil servant. Under the parallel trend assumption, the treatment and control group should not display pronounced differential time patterns in absence of treatment. Table 2 shows that the coefficients on the interactions of dummies for year and the husband being employed as a civil servant are not significant for the pretreatment years 2000–2003. This suggests that the parallel trend assumption holds as far as the pretreatment years are concerned.

Adding a vector of covariates z_{it} as displayed in Table 1 and dummy variables for the wife's

12. Recall that we treat retirement of the first retiring member of the couple as an absorbing state. Hence, H_{it} can only be 1 if the husband retired before the wife.

13. The parameters are α_H (intercept), β_{jH} (year fixed effects), γ_{jH}(year j times employed as a civil servant), and v_{it} is an error term.

TABLE 2
Husband's Probability to Retire: Testing the Common Trend Assumption

	Coefficient	Std. Err.	*p* Value
Husband civil servant × Year 2005	0.0819	0.0065	0.000
Year 2000	−0.0023	0.0032	0.480
Year 2001	0.0032	0.0030	0.288
Year 2002	0.0096	0.0028	0.001
Year 2003	−0.0045	0.0026	0.080
Year 2004	−0.0023	0.0025	0.370
Husband civil servant	−0.0214	0.0036	0.000
Husband civil servant × Year 2000	0.0091	0.0071	0.196
Husband civil servant × Year 2001	0.0063	0.0065	0.329
Husband civil servant × Year 2002	−0.0079	0.0057	0.164
Husband civil servant × Year 2003	0.0010	0.0052	0.843
Constant	0.0589	0.0018	0.000
N	95,141		

Notes: This table shows OLS estimates of a linear probability model of husband's retirement status within 1 year (12 months). There are no other regressors. Estimates are based on the baseline sample (Table 1). Treatment group: Husband civil servant; treatment year: 2005. Standard errors are robust to heteroscedasticity.

age A_{kitW} (taking the value 1 if the wife's age in year t is equal to k), and removing dummy interactions of pretreatment years with the husband being employed as a civil servant, we perform a difference-in-difference analysis with a treatment group that is homogeneous in age (55–60):

$$(2) \qquad H_{it} = \alpha_H + \sum_{j=2000}^{j=2004} \beta_{jH}\iota(t=j) + \gamma_{2005,H}\iota(t=2005)\,C_{itH} + \zeta_H C_{itH} + \sum_{k=32}^{k=66} \chi_{kW} A_{kitW} + z'_{it}\varphi_W + v_{it}.$$

The coefficient of interest, $\gamma_{2005,\,H}$, measuring the interaction of year 2005 and the treatment group, has a point estimate of 9 percentage points and is highly significant. Table 3 reports full results on the exercise.

C. Intention-to-Treat Estimates

We perform a similar difference-in-difference analysis for wife's retirement status to study the spillovers of husbands' induced retirement on their wives' retirement behavior.

We again start with discussing whether the parallel trend assumption (1) and the assumption on the absence of compositional changes (2) hold. Figure 2 shows that prereform trends in wife's retirement rates were parallel for the control and treatment group. Wife's retirement rates show a jump of about 1 percentage point in 2005 for the treatment group. We also formally verify the parallel trend assumption by estimating Equation (1) with the wife's retirement status as the dependent variable. Table 4 shows that none of the coefficients on the interactions of the dummies for the pretreatment years 2000–2003 and the husband being employed as a civil servant are significant.

Concerning the second assumption, we discussed in Section IV that there were no large differences in observables between the control and treatment group before and after the reform. We also have no reason to believe that such differences in unobservables exist.

Spillover effects of husbands' retirement incentives on their wives' retirement behavior can be captured through specifying the following equation:

$$(3) \quad W_{it} = \alpha_W + \sum_{j=2000}^{j=2004} \beta_{jW}\iota(t=j) + \gamma_{2005,W}\iota(t=2005)\,C_{itH} + \zeta_W C_{itH} + \sum_{k=32}^{k=66} \chi_{kW} A_{kitW} + z'_{it}\varphi_W + u_{it},$$

where W_{it} is a dummy variable that is 1 if the wife in couple i retired in year t and 0 otherwise, and u_{it} is the error term. All other (year) variables are as specified above for Equation (2). Coefficient $\gamma_{2005,\,W}$ in particular, measures the ITTE, that is, the spillover effect of the policy change (directly affecting the husband) on the retirement behavior of the wife.

Table 5 shows results. The ITTE coefficient estimate indicates that the policy change increased the probability of wives' retirement whose husbands were employed as civil servants by 1 percentage point. The effect is significant at the 5% level, but it is an order of magnitude smaller than the effect for husbands.

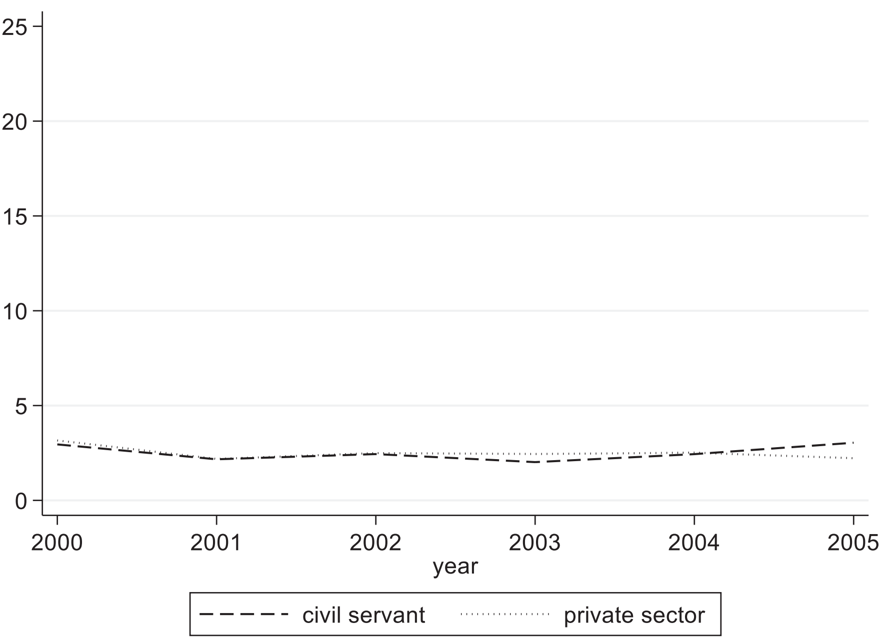
Further refinement by splitting the treatment group in two (ages 55–57 vs. 58–60) is motivated by the greater strength of the retirement incentive for older ages. Figure 3 shows that husband's retirement rates for the two subgroups display very different behavior. The upper age

TABLE 3
Husband's Probability to Retire: Difference-in-Difference Estimates

	Coefficient	Std. Err.	p Value
Husband civil servant × Year 2005	0.0879	0.0069	0.000
Wife's wage income [$t-1$] (in 10,000s of Euros)	0.0009	0.0006	0.104
Husband's wage income [$t-1$] (in 10,000s of Euros)	−0.0026	0.0003	0.000
Wife hospitalized [$t-1$]	−0.0002	0.0039	0.949
Husband hospitalized [$t-1$]	0.0105	0.0031	0.001
Wife Dutch nationality	0.0023	0.0023	0.321
Husband Dutch nationality	0.0002	0.0023	0.923
Wife's age dummies	Yes		
Year dummies	Yes		
Husband civil servant	−0.0211	0.0020	0.000
Constant	0.1284	0.0414	0.002
N	95,141		

Notes: This table shows OLS estimates of a linear probability model of husband's retirement status within 1 year (12 months). Estimates are based on the baseline sample (Table 1). Treatment group: Husband civil servant; treatment year: 2005. Standard errors are robust to heteroscedasticity.

FIGURE 2
Wives' Retirement Rates by Year and Sector (Percentages)



Notes: This figure shows average 12-months-ahead retirement rates for wives, by husbands' sector of employment. Based on the baseline sample (Table 1).

group shows a jump of about 16% and the lower age group a jump of about 3% (the total effect is a weighted average with the lower age group being more populous). We skip displaying and discussing estimation results corresponding to Equation (2) for this refinement (the results can be obtained on request and are fully in line with Figure 3), and instead focus directly on the extension of Equation (3), modeling retirement decisions of the wife.

Table 6 shows that if we split the treatment group in two (55–57 and 58–60), the ITTE coefficient estimate for wives is not significant for the treatment group having husbands in the age

TABLE 4
Wife's Probability to Retire: Testing the Common Trend Assumption

	Coefficient	Std. Err.	*p* Value
Husband civil servant × Year 2005	0.0090	0.0027	0.001
Year 2000	0.0093	0.0016	0.000
Year 2001	−0.0004	0.0012	0.736
Year 2002	0.0027	0.0011	0.014
Year 2003	0.0022	0.0010	0.027
Year 2004	0.0030	0.0009	0.001
Husband civil servant	−0.0008	0.0009	0.345
Husband civil servant × Year 2000	−0.0013	0.0033	0.693
Husband civil servant × Year 2001	0.0006	0.0026	0.817
Husband civil servant × Year 2002	0.0003	0.0023	0.895
Husband civil servant × Year 2003	−0.0035	0.0022	0.114
Constant	0.0223	0.0147	0.130
N	95,141		

Notes: This table shows OLS estimates of a linear probability model of wife's retirement status within 1 year (12 months). There are no other regressors. Estimates are based on the baseline sample (Table 1). Treatment group: Husband civil servant; treatment year: 2005. Standard errors are robust to heteroscedasticity.

TABLE 5
Wife's Probability to Retire: Intention-to-Treat Effect

	Coefficient	Std. Err.	*p* Value
Husband civil servant × Year 2005	0.0099	0.0032	0.002
Wife's wage income [$t-1$] (in 10,000s of Euros)	−0.0008	0.0004	0.067
Husband's wage income [$t-1$] (in 10,000s of Euros)	0.0010	0.0002	0.000
Wife hospitalized [$t-1$]	0.0119	0.0032	0.000
Husband hospitalized [$t-1$]	0.0015	0.0019	0.454
Wife Dutch nationality	−0.0035	0.0016	0.027
Husband Dutch nationality	−0.0043	0.0016	0.007
Wife's age dummies	Yes		
Year dummies	Yes		
Husband civil servant	−0.0023	0.0015	0.115
Constant	0.0134	0.0136	0.323
N	95,141		

Notes: This table shows OLS estimates of a linear probability model of wife's retirement status within 1 year (12 months). Estimates are based on the baseline sample (Table 1). Treatment group: Husband civil servant; treatment year: 2005. Standard errors are robust to heteroscedasticity.

category 55–57; the estimate is a noticeable and significant 2 percentage points for the treatment group with husbands in the age category 58–60. As before, the control group consists of wives with husbands employed in the private sector and in the age category 55–60.

D. *Instrumental Variable Analysis*

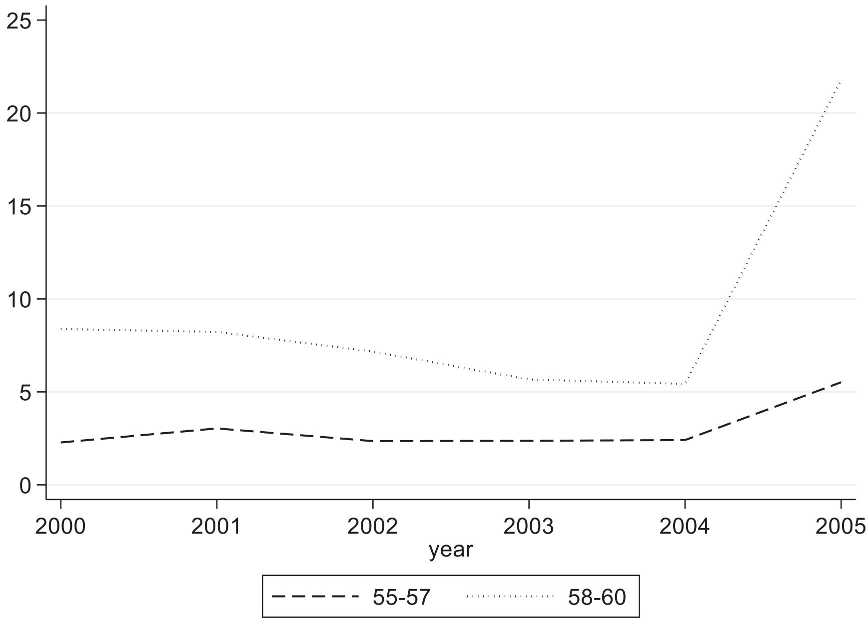
Given that husbands and wives can retire from the labor market at virtually any age and year, for a large variety of reasons (ranging from income effects to health considerations), the interesting question we want to answer is now how much of an effect in wives' ER can be causally ascribed to the changed-incentive-induced ER of husbands (LATE)? We apply an IV approach (two stage least squares), instrumenting the husband's retirement status with his treatment group membership (civil servant) interacted with age dummies and treatment year 2005. So the first stage is a refinement of the difference-in-difference analysis where we used two eligible age groups. In the second stage, the impact of predicted retirement of the husband on the wife's probability to retire within 12 months is estimated. The reduced form corresponds to a (further refined) difference-in-differences analysis for the wife's retirement. As the analysis (Section V.C, Table 6) showed, there is strong evidence for causal spillover effects of retirement status between spouses, and there is reason to assume that treatment effects are heterogeneous across age groups. The finer-grained split of age groups can help us understand who, within the treatment group, is affected by how much.

We specify the first stage of our model as follows:

$$(2') \; H_{it} = \alpha_H + \sum_{j=2000}^{j=2004} \beta_{jH} \iota (t = j) + \sum_{k=32}^{k=66} \chi_{kH} A_{kitW} + \sum_{l=56}^{l=60} \delta_{lH} A_{litH} + \sum_{l=55}^{l=60} \eta_{lH} A_{litH} C_{itH} + \sum_{l=55}^{l=60} \kappa_{lH} \iota (t = 2005) \times A_{litH} C_{itH} + z'_{it} \varphi_H + v_{it}.$$

The specification is a generalization of Equation (2) that formed the basis for our difference-in-difference analysis, now controlling for age-specific treatment effects, spousal ages, and additional regressors. Equation (2′)

FIGURE 3
Husbands' Retirement Rates by Year and Sector (Civil Servants; Percentages)



Notes: This figure shows average 12-months-ahead retirement rates for husbands employed as civil servants, by husbands' age group. Selected from the baseline sample (Table 1, panels b and d).

contains, subsequently, year dummies, cross effects year dummies, and civil servants, wife's age dummies, husband's age dummies, cross effects of husband's age dummies and civil servants, the instruments, that is, age-specific treatment effects (for civil servants in 2005), and additional covariates.

The second stage is specified as follows:

$$(3') \; W_{it} = \alpha_W + \sum_{j=2000}^{j=2004} \beta_{jW} \iota(t=j) + \sum_{k=32}^{k=66} \chi_{kW} A_{kitW} + \sum_{l=56}^{l=60} \delta_{lW} A_{litH} + \sum_{l=55}^{l=60} \eta_{lW} A_{litH} C_{itH} + z'_{it}\varphi_W + \omega \widehat{H}_{it} + u_{it}.$$

Again, the equation is similar to Equation (3), except that the impact of the policy variables is not directly incorporated but will be reflected through the husband's predicted retirement choice. ω, the coefficient on the predicted retirement indicator of the husband, indicates the LATE. u_{it} and v_{it} are allowed to be correlated with each other.

Instrument Validity. The validity of our instruments hinges on the satisfaction of two conditions. First, the instruments have an impact on the probability that husbands retire. Second, the instruments do not correlate with unobserved factors having an impact on the probability that wives retire.

Figure 4 shows retirement rates for different birth cohorts by age for husbands who were employed as civil servants. As we saw before in Figures 1 and 3, husbands had higher retirement rates in 2005 than in earlier years. The difference in retirement rates between 2005 and earlier years varies by age and was especially large for husbands in the upper age category 58–60, considerably smaller for husbands aged 57, and even smaller for husbands aged 55–56. This is all in line with the age-specific incentives as provided by the temporary decrease in eligibility age in ER benefits, as discussed in Section III. This supports our hypothesis that our instruments are relevant.

To our knowledge, there were in 2005 no similar ER windows in sectors other than the public sector. We thus do not expect the opening of the ER window to have had a direct impact on the probability that the wives retired. In Appendix

TABLE 6
Wife's Probability to Retire, Heterogeneous Intention-to-Treat Effect

	Coefficient	Std. Err.	p Value
Husband aged 55–57 × Husband civil servant × Year 2005	0.0055	0.0032	0.084
Husband aged 58–60 × Husband civil servant × Year 2005	0.0168	0.0061	0.006
Wife's wage income [$t-1$] (in 10,000s of Euros)	−0.0009	0.0004	0.052
Husband's wage income [$t-1$] (in 10,000s of Euros)	0.0010	0.0002	0.000
Wife hospitalized [$t-1$]	0.0119	0.0032	0.000
Husband hospitalized [$t-1$]	0.0013	0.0019	0.500
Wife Dutch nationality	−0.0034	0.0016	0.035
Husband Dutch nationality	−0.0043	0.0016	0.007
Wife's age dummies	Yes		
Husband aged 58–60	0.0052	0.0013	0.000
Husband aged 55–57 × Husband civil servant	−0.0020	0.0015	0.186
Husband aged 58–60 × Husband civil servant	−0.0028	0.0031	0.362
Year dummies	Yes		
Constant	0.0112	0.0135	0.406
N	95,141		

Notes: This table shows OLS estimates of a linear probability model of wife's retirement status within 1-year (12 months). Estimates are based on the baseline sample (Table 1). Treatment group: Husband civil servant in the age category 55–60; treatment year: 2005. Standard errors are robust to heteroscedasticity.

B2, we perform a robustness check using placebo instruments for the eligibility of private sector workers. We are also not aware of any event other than the opening of the ER window that may have affected the probability to retire for civil servants in 2005. We expect that our instruments are not correlated with unobserved factors that influenced the wives' probability to retire. Wives' unobserved health, number of hours worked, or stress levels associated with work are among the unobserved factors that may have influenced wives' probability to retire. These factors may be affected if retirement of the husband could have been anticipated. Correlation due to anticipation is not expected to be an issue. This is because the opening of the ER window was only announced by the government in April 2004 and central government employers only decided within the next 8–9 months whether and to whom they would actually open the ER window. Selection into public sector jobs after the announcement of the policy change is not an issue either, because eligibility was tied to career qualifications described earlier.

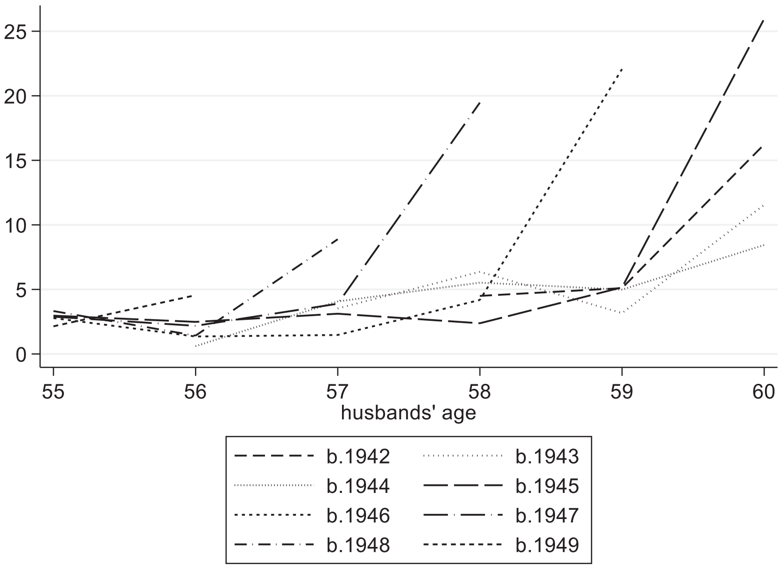
Results for the Uninstrumented Case. For comparison, we first estimate by ordinary least squares (OLS) the model as specified in (3′) without correcting for potential endogeneity of husband's retirement, so with husband's retirement status rather than husband's predicted retirement status as the key regressor. Table 7 shows that the coefficient estimate of husband's actual retirement status on the wife's retirement is (less than) 1 percentage point and significant at the 5% level. This effect is similar in size to the ITTE.

Instrumental Variable Estimates. Table 8 shows the IV estimate. The coefficient estimate on retirement of the husband indicates that retirement of husbands induced by the opening of the ER window actually increased their wives' probability to retire within 1 year by 10 percentage points. This effect is significant at the 5% level. Interestingly, the coefficient estimate on retirement of the husband is larger than it was in the uninstrumented case. One explanation for the LATE being larger is that those likely to respond to the policy were those who had wives who would retire with them. It may also be that the difference in coefficient estimates is due to endogeneity bias in the uninstrumented case.[14] The F-statistic in the first stage shows that our instruments are jointly significant at the 1% level. The LATE is considerably larger than the ITTE discussed above, as can be expected since not all treated husbands actually do retire. The LATE estimate is of a similar magnitude as the effect found by Banks, Blundell, and Casanova Rivas (2010). They find a positive effect of 14–20 percentage points of incentive-induced retirement of the husband on wives' probability to retire within 2-year time intervals for husbands that are between zero and 2 years older than their wives.

We should point out at this stage that the finding of a 10 percentage point is fully robust to alternative identification strategies and changes in the specification of the conditional mean

14. The coefficient on retirement status of the husband can be biased in both directions through many different mechanisms. This also makes it difficult to indicate which mechanism or mechanisms made the coefficient biased downwards rather than upwards.

FIGURE 4
Husbands' Retirement Rates by Age and Year of Birth (Civil Servants; Percentages)



Notes: This figure shows average 12-months-ahead retirement rates for husbands employed as civil servants, by husbands' birth cohort. Selected from the baseline sample (Table 1, panels b and d). 2005 minus husbands' year of birth indicates husbands' age in treatment year 2005.

function. In Appendix B1, we use workers employed as civil servants and in the (ineligible) age category 52–54 in 2005 as an alternative control group. We use two samples, one with observations on wives with husbands employed as civil servants only, and another one also including observations on wives with husbands employed in the private sector. We arrive at very similar point estimates for the causal effect. Further changes in the functional form specification are investigated in Appendix B3.

Reduced Form Estimates. The coefficient estimates on the instruments in the reduced form model can show us how old the husbands are that drive the effect of incentive-induced husband's retirement on the wife's probability to retire within 1 year. The reduced form of model (3′) (resulting from inserting (2′) in (3′)) actually is a refined variant of the difference-in-difference specification (3), with separate treatment coefficients for each eligible age. Table 9 shows the reduced form estimates. In the earlier analysis based on (3), with all eligible ages pooled, we found a positive and significant effect of treatment (Table 5). The further refinement (Table 9) shows that the treatment coefficients of each

TABLE 7
Wife's Probability to Retire: Effect of the Husband's Retirement Status

	Retirement Status of the Wife		
	Coefficient	**Std. Err.**	***p* Value**
Retirement status husband	0.0117	0.0028	0.000
Wife's wage income [$t-1$] (in 10,000s of Euros)	−0.0009	0.0004	0.043
Husband's wage income [$t-1$] (in 10,000s of Euros)	0.0010	0.0002	0.000
Wife hospitalized [$t-1$]	0.0118	0.0032	0.000
Husband hospitalized [$t-1$]	0.0012	0.0019	0.538
Wife Dutch nationality	−0.0033	0.0016	0.037
Husband Dutch nationality	−0.0042	0.0016	0.008
Wife's age dummies	Yes		
Husband's age dummies	Yes		
Husband's age dummies × Husband civil servant	Yes		
Year dummies	Yes		
Constant	0.0106	0.0137	0.439
N	95,141		

Notes: This table shows OLS estimates of a linear probability model of wife's retirement status within 1 year (12 months), as function of the observed retirement status of the husband. Estimates are based on the baseline sample (Table 1). Standard errors are robust to heteroscedasticity.

TABLE 8

Wife's Probability to Retire (Instrumental Variables)

	First Stage Retirement Status of the Husband			Second Stage Retirement Status of the Wife		
	Coefficient	**Std. Err.**	***p* Value**	**Coefficient**	**Std. Err.**	***p* Value**
Predicted retirement status husband				0.1042	0.0372	0.005
Instruments (6x)						
Husband's age 55 × Husband civil servant × Year 2005	0.0102	0.0072	0.156			
Husband's age 56 × Husband civil servant × Year 2005	0.0325	0.0078	0.000			
Husband's age 57 × Husband civil servant × Year 2005	0.0609	0.0111	0.000			
Husband's age 58 × Husband civil servant × Year 2005	0.1537	0.0164	0.000			
Husband's age 59 × Husband civil servant × Year 2005	0.1719	0.0177	0.000			
Husband's age 60 × Husband civil servant × Year 2005	0.1386	0.0268	0.000			
Wife's wage income [$t-1$] (in 10,000s of Euros)	0.0009	0.0006	0.099	−0.0010	0.0004	0.029
Husband's wage income [$t-1$] (in 10,000s of Euros)	−0.0026	0.0003	0.000	0.0013	0.0003	0.000
Wife hospitalized [$t-1$]	−0.0003	0.0039	0.930	0.0119	0.0032	0.000
Husband hospitalized [$t-1$]	0.0105	0.0031	0.001	0.0002	0.0020	0.906
Wife Dutch nationality	0.0023	0.0023	0.321	−0.0035	0.0016	0.028
Husband Dutch nationality	0.0002	0.0023	0.915	−0.0042	0.0016	0.008
Wife's and husband's age dummies, year dummies	Yes			Yes		
Husband's age dummies × Husband civil servant	Yes			Yes		
Constant	0.1245	0.0414	0.003	−0.0024	0.0150	0.874
F statistic instruments	40.16					
N	95,141					

Notes: This table shows 2SLS estimates of a linear probability model of wife's retirement status within 1 year (12 months), as function of the predicted retirement status of the husband. Estimates are based on the baseline sample (Table 1). Standard errors are robust to heteroscedasticity.

TABLE 9

Wife's Probability to Retire (Reduced Form)

	Coefficient	**Std. Err.**	***p* Value**
Instruments (6x)			
Husband's age 55 × Husband civil servant × Year 2005	0.0006	0.0042	0.877
Husband's age 56 × Husband civil servant × Year 2005	0.0097	0.0053	0.067
Husband's age 57 × Husband civil servant × Year 2005	0.0055	0.0063	0.386
Husband's age 58 × Husband civil servant × Year 2005	0.0040	0.0074	0.592
Husband's age 59 × Husband civil servant × Year 2005	0.0149	0.0098	0.130
Husband's age 60 × Husband civil servant × Year 2005	0.0465	0.0174	0.007
Wife's wage income [$t-1$] (in 10,000s of Euros)	−0.0009	0.0004	0.045
Husband's wage income [$t-1$] (in 10,000s of Euros)	0.0010	0.0002	0.000
Wife hospitalized [$t-1$]	0.0118	0.0032	0.000
Husband hospitalized [$t-1$]	0.0013	0.0019	0.499
Wife Dutch nationality	−0.0033	0.0016	0.039
Husband Dutch nationality	−0.0042	0.0016	0.008
Wife's and husband's age dummies, year dummies	Yes		
Husband's age dummies × Husband civil servant	Yes		
Constant	0.0102	0.0135	0.452
F statistic instruments	2.18		
N	95,141		

Notes: This table shows OLS estimates of a linear probability model of wife's retirement status within 1 year (12 months), as function of the instruments. Estimates are based on the baseline sample (Table 1). Standard errors are robust to heteroscedasticity.

eligible age separately are not all individually significant at the 5% level. The coefficient on husband's age 60 interacted with the dummy for the year 2005 is the only one that is positive and significant, indicating that the effect of interest is driven by wives with husbands aged 60. As husbands were typically older than or as old as their wives, husbands aged 60 have on average the highest likelihood of having wives aged 60. In the observation period, the age of 60 was a

TABLE 10
Wife's Probability to Retire (Alternative Reduced Form)

	Coefficient	Std. Err.	p Value
Selected coefficients			
Wife's age 55 × Husband civil servant × Year 2005	0.0110	0.0079	0.162
Wife's age 56 × Husband civil servant × Year 2005	0.0024	0.0074	0.747
Wife's age 57 × Husband civil servant × Year 2005	0.0101	0.0102	0.324
Wife's age 58 × Husband civil servant × Year 2005	0.0150	0.0152	0.323
Wife's age 59 × Husband civil servant × Year 2005	0.0336	0.0202	0.097
Wife's age 60 × Husband civil servant × Year 2005	0.1437	0.0543	0.008
Wife's age 61 × Husband civil servant × Year 2005	−0.0255	0.0773	0.742
Wife's age 62 × Husband civil servant × Year 2005	0.1805	0.1422	0.204
Wife's age 63 × Husband civil servant × Year 2005	0.3545	0.3522	0.314
Wife's age dummies × Husband civil servant × Year 2005	Yes		
Wife's wage income [$t-1$] (in 10,000s of Euros)	−0.0009	0.0004	0.045
Husband's wage income [$t-1$] (in 10,000s of Euros)	0.0010	0.0002	0.000
Wife hospitalized [$t-1$]	0.0118	0.0032	0.000
Husband hospitalized [$t-1$]	0.0012	0.0019	0.543
Wife Dutch nationality	−0.0033	0.0016	0.038
Husband Dutch nationality	−0.0042	0.0016	0.009
Wife's and husband's age dummies, year dummies	Yes		
Husband's age dummies × Husband civil servant	Yes		
Constant	−0.0026	0.0028	0.358
N	95,141		

Note: This table shows OLS estimates of a linear probability model of wife's retirement status within 1 year (12 months), as function of (interactions of) public/private sector group indicators and wife's age indicators and treatment year indicators. Estimates are based on the baseline sample (Table 1). Standard errors are robust to heteroscedasticity.

common age of eligibility for ER in many private sector ER arrangements. This is for us a reason to verify whether the effect of retirement of husbands on retirement status of wives runs through husbands with wives aged 60.

Table 10 shows the coefficient estimates of a linear regression model with wife's retirement status as the dependent variable and wife's age dummies interacted with the dummy for the year 2005.[15] The coefficients on all of those interaction terms are insignificant, except the one for the interactions between age 60 and year 2005. That coefficient is large and positive. This indicates that wives aged 60 in 2005 had a much larger probability to retire within 1 year than wives aged 60 in earlier years. Banks, Blundell, and Casanova Rivas (2010) also find such an interaction effect of retirement incentives.

VI. CONCLUSIONS

We identify and estimate the causal effect of the husband's retirement status on the wife's retirement rate by using a quasi-natural experiment where policy variation induces the husband to retire early. We focus on couples where only the male, but not the female, was subject to specific ER incentives applicable to civil servants, and both spouses had a strong labor market attachment.

We benefit from having administrative data from the Netherlands and having a strong and exogenous source of variation in retirement status of the husband at our disposal. We find that incentive-induced retirement of husbands increases the probability that the wife retires within 12 months by 10 percentage points. The finding indicates that the temporary decrease of the eligibility age for ER benefits for male civil servants has a strong (indirect, but nearly immediate) effect on retirement status of their spouse. Our result is robust to changes in functional form specification, and in general robust to changes in data selection criteria as well. It is also robust to employing alternative identification strategies (identifying assumptions and definition of control groups). All in all, we believe that leisure complementarities in preferences are consistent with the pattern documented here. However, we find a strong effect coming from a precisely characterized group, namely husbands aged 60 with wives aged 60. As age 60 was the eligibility age for regular ER benefits in many sectors, our result

15. Also included in the regressor set are wife's age dummies, husband's age dummies, husband's age dummies interacted with the dummy for the husband being employed as a civil servant, year dummies, and wife's and husband's personal characteristics.

suggests that retirement of husbands induced by the opening of the ER window may have caused wives to retire using regular ER benefits. Alternatively, husbands with wives aged 60 may have been more likely to accept the ER offer than husbands with wives of other ages. Such interaction effects of the ER window under review and regular ER arrangements in other sectors would be consistent with the evidence for interaction effects for ER arrangements for spouses provided by Banks, Blundell, and Casanova Rivas (2010).

Reforms are ubiquitous in retirement systems around the world, and there is much experimentation with policy parameters, sometimes instigated by idiosyncratic ad-hoc decisions. The reform we study is no exception, as it advanced retirement entry age in a world characterized by tightening qualification or abolition of ER. We may speculatively conclude that many policy initiatives that postpone final labor market withdrawal likewise may have a double impact—not only on the workers directly affected, but indirectly by way of delaying retirement of their spouses.

APPENDIX A: INSTITUTIONAL SETTINGS

A1. THE DUTCH PENSION AND RETIREMENT SYSTEM

The Dutch pension system rests on three pillars (Bovenberg and Meijdam 2001).[16] The first pillar is the public old-age pension system (social security), financed on a pay-as-you-go basis. Contributions stem from workers and employers. All residents registered in the Netherlands accrue public old-age pension rights. Public old-age pension benefits are flat. For couples they equal the minimum wage. Singles receive 70% of the minimum wage. For every year between the ages 15 and 65 that an individual does not reside in the Netherlands, public old-age benefits are cut by 2 percentage points. The second pillar consists of occupational pensions (including company-specific funds of large firms, and industry-wide funds covering all occupations in an industry). These are funded pensions, typically of the defined benefit type, and generally managed on the sector (or firm) level. The third pillar consists of private provisions. Those include, among others, annuities or life insurance policies.

About 90% of all employees participate in an occupational pension plan. Occupational pension schemes receive contributions from workers and employers. Workers who participate in a pension plan pay contributions over the difference between their wage and a nominal threshold called the “franchise.” The franchise is about 143% of the public-old age pension benefit level. In this way, first-pillar old age pensions and second-pillar occupational pensions are integrated and jointly achieve before-tax replacement rates of 70%. As every sector has its own pension plan and pension conditions, there is a large heterogeneity among occupational pensions.

16. See also Bloemen, Hochguertel, and Zweerink (2017).

A2. REGULAR EARLY RETIREMENT ARRANGEMENTS FOR PRIVATE SECTOR WORKERS

For the time period studied, early retirement (ER) pensions, embedded in the occupational pension system, were widespread, owing to tax incentives.[17] There was an accordingly large heterogeneity in ER arrangements across sectors. The ER eligibility age generally varied from age 60 to 62. Workers and employers contributed to both parts of the ER benefit scheme via a separate pay-as-you-go system. Gross (before-tax) benefits corresponded to 70% of the preretirement average salary level. Retiring through the ER system had no effect on regular pension benefits (for which eligibility started at age 65).

A3. REGULAR EARLY RETIREMENT ARRANGEMENTS FOR CIVIL SERVANTS

The public sector had its own ER system, which shared many features with the private sector schemes, however. As of April 1, 1997, ER benefits for civil servants consisted of two parts. The first part was financed on a pay-as-you-go basis and was in general 70% of the franchise for civil servants who had worked full-time during their working life.[18] The second part was funded and complemented the first part to a sum of up to 70% of workers’ average pay (mid-career salary). Workers and employers contributed to both parts of the ER benefit scheme. The first part intended to compensate early retirees for the lack of old-age pension benefits for the period between ER and normal retirement. Civil servants were eligible for the first part if they satisfied two conditions. First, they had to have been employed as civil servants continuously during the 10 years prior to ER. Second, they had to have contributed continuously to the public pension fund during the 10 years preceding ER. The first part of ER benefits was in general higher when a civil servant retired at a later age. ER among civil servants usually occurred at age 61 or 62.

Pensions achieving the maximum replacement rate required a contribution history of 40 years. The replacement rate was reduced by 1.75 percentage points for every year that the total pension contribution period fell short of 40 years. Civil servants were allowed to do paid work after ER. However, total income of a retired civil servant was not allowed to exceed 100% of the average pay. Otherwise, ER benefits were cut to bring the total income earned to 100% of the average gross wage (means test).

A4. THE 2005 ER WINDOW

Following announcement in April 2004, departments (ministries) could offer their qualifying employees the possibility to retire as early as of age 55 during the months of January through November 2005. The benefit duration was capped at 8 years, however. Qualification entailed meeting

17. The so-called fiscal facilitation of the early retirement contributions implied that the early retirement benefits were taxed, and that the early retirement contributions paid by workers and employers were exempted from taxation. As effectively less tax was paid, the fiscal facilitation made early retirement very attractive for both eligible workers and employers.

18. This replacement rate is based on retirement at the ER eligibility age.

TABLE B1
Robustness Checks on Data Selection (IV Estimates)

Variation	LATE	Std. Err.	p Value	F Stat.	N
Baseline (Table 8)	0.1042	0.0372	0.005	40.16	95,141
a. Sample without restricting wage income [$t-1$]	0.0286	0.0309	0.354	86.03	246,351
b. Sample incl. husbands aged 52–60	0.0983	0.0365	0.007	41.84	170,619
c. Sample incl. husbands aged 52–60, civil servants only	0.0904	0.0394	0.022	42.60	30,344

Note: See Table 8 for estimator, specification, and baseline sample. Variations alter, one by one, the data selection criteria.

the 10-year employment and contribution rules that applied to regular ER arrangements as well.

We exemplify the different incentives operating for civil servants at different ages in 2005. Those, say, born in January 1950, who just turned 55 and retired in January 2005, would see their ER benefits expire in January 2013 at the age of 63. They would then need to bridge another 2 years without benefits, before becoming eligible for regular pension benefits in January 2015 again. Those born in January 1948, who turned 57 in January 2005, could have retired on the special ER scheme and continued on regular pension benefits, not facing a coverage gap. Those born before January 1948 could in addition continue accruing regular pension claims at a rate of 50% at their employer's expense for up to 4 years.

A5. OTHER POLICY CHANGES

On January 1, 2004, the public sector pension fund switched from a final pay pension regime to an average pay pension regime. Owing to a transitional arrangement, civil servants born before January 1, 1954 were hardly affected by this switch.

On January 1, 2006, the so-called fiscal facilitation of ER benefits for individuals born on January 1, 1950 or later was terminated. This implied that most ER arrangements for the youngest workers in our sample disappeared. The factual abolition of ER arrangements may have induced the affected workers to retire later. The termination of the fiscal facilitation of ER benefits could have been anticipated as of 2003 and may have induced anticipation effects of civil servants aged 53–55 in 2005.

A6. EARLY RETIREMENT VIA DI

Workers may alternatively effectuate an early withdrawal from the labor force by starting to receive DI benefits. Workers could start receiving DI benefits if they would be judged to be disabled by a medical examiner of the social insurance institute. DI benefits amounted to up to 70% of the final pay. The DI system was financed by workers and employers. In the context of this paper, it is important to notice that civil servants faced the same generosity of and eligibility criteria for DI benefits as workers employed in other sectors. Hence, program substitution effects will not explain our findings.

There were also no major changes in DI in 2005 that could potentially explain our findings. There was one change in DI implemented on December 29, 2005. From that day on, eligibility criteria for starting to receive disability benefits have been tightened and disability benefits have been made less generous for workers who were only partially disabled. One year earlier, another change in DI had taken place. Since January 1, 2004, workers could only start receiving DI benefits after having been on continuous sick leave for 2 years. We hardly observe any civil servants who were induced to retire in 2005 and who started receiving disability benefits soon after retirement. Moreover, the inflow into DI for men in all ages decreased smoothly from 53,000 in 2001 to 36,000 in 2004 and dropped to 17,000 in 2005. It is thus unlikely that the retirement rates in 2005 were boosted by changed incentives in DI (Statistics Netherlands 2006).

APPENDIX B: ROBUSTNESS CHECKS

B1. ROBUSTNESS CHECK ON DATA SELECTION

B.1.1 No selection on wage income

We selected observations on workers who had a lagged wage income of at least 20,000 euros. Not having recourse to information on hours worked, we did so to make sure that husbands and wives in our sample had a strong labor force attachment. Workers with a weak labor force attachment may not have performed career planning and may not have planned or coordinated retirement with their spouses. Also, workers with low wage income and low hours worked may have had much leisure time already, so leisure complementarities may have been less of a motivation for them to coordinate retirement. We may therefore not expect an effect for this group of workers. Table B1 shows that the LATE estimate is indeed not significant if observations on workers with a wage income (at $t-1$) of any level are included (variation a). Note that the first stage continues to be strong (F-test), meaning that husbands were kept being induced to retire. Hence, we interpret this as evidence for the hypothesis that women with a stronger labor force attachment were indeed more likely to coordinate retirement with their spouses.

B.1.2 Alternative identification strategies (control groups)

Our baseline sample has couples in which husbands are in the age category 55–60 only. Alternatively, we can add observations on couples in which husbands were in the age category 52–54 (and working in either sector) to the sample. Note that all of the added husbands were in the control group (the added public sector workers were ineligible for the ER window).

The reason for this alternative setup is that not only the nature of labor contracts (including employment protection) differs substantially between private and public sectors, but also ER arrangements, while similar, differ across sectors in many details. So, there may be unobservable, but relevant differences between the treatment and control group that our baseline specification fails to pick up. The disadvantage of the present alternative identification strategy is that workers in the control group are not necessarily in the same age category as those in the treatment group (the younger ones in fact have very low retirement rates).

We find, however, that the LATE estimated using the alternative model (variation b) is similar to the LATE estimated using the baseline model.

We can even go a step further and drop observations on couples in which husbands were employed in the private sector. The advantage of this approach is that results are not affected by possibly unnoticed developments in retirement rates in the private sector. Identification relies therefore on age in 2005.

The LATE estimated using the alternative model is, however, again similar to our baseline LATE (variation c). Both these two alternative identification strategies, in conjunction with the baseline result, therefore provide reliable evidence for the validity of our approach and robustness of our finding.

B2. PLACEBO INSTRUMENTS

We further verify whether the variation in retirement rates of husbands in the age category 55–60 in 2005 could have been due to another event that occurred in 2005 and may have affected the private sector workers as well.

We deviate from the baseline by choosing as instruments dummies for the husband's age (55–60) interacted with a dummy for the year 2005 interacted with a dummy for the husband being employed in the private sector. We also include an interaction of dummies for the year 2005 and the husband being employed as a civil servant to absorb the effect of the treatment received by civil servants.

Table B2 shows that there is no effect of placebo incentive-induced retirement of the husband on retirement status of the wife. The first stage coefficient estimates on the instruments show that the interaction of husband's age 55 with the year 2005 dummy and the husband being employed in the private sector dummy is the only instrument that significantly affects retirement status of the husband. That coefficient is only small in size. This suggests that the jumps in retirement rates for husbands employed as civil servants are induced by the opening of the ER window rather than a 2005 macro shock that shifted the retirement rates of all employed husbands upwards.

B3. ROBUSTNESS CHECKS ON THE MODEL SPECIFICATION

B.3.1 Wife's retirement horizon

Our baseline IV model uses wife's retirement within 12 months as the dependent variable in the second stage. However, our results may be sensitive to the time horizon for which we measure wife's retirement. The wife's reaction to the decision of her husband may be somewhat delayed, for instance, for many reasons. Table B3 shows that the LATE estimate is similar to the baseline estimate if we use wife's retirement within 24 months as the dependent variable (variation a). This indicates that the effect of induced husband's retirement on wife's retirement is rather immediate and concentrated on wife's retirement within 12 months. The proximity in time of the spousal retirement decisions bolsters the hypothesis of leisure complementarities that may be at work even for individuals with a strong labor market attachment.

B.3.2 Industry dummies

Husbands and wives employed in the private sector are heterogeneous groups of workers, possibly more so than in the public sector. Our baseline specification does not control for the heterogeneity of retirement rates across industries. Adding industry dummies for both wives and husbands does not affect our results, however (variation b).[19]

B.3.3 Exclusion of wage income and hospitalization

Table 1 suggested that there are small but occasionally significant differences across treatment/control groups for some of the regressor means. It may therefore be important to control for those regressors explicitly, as we do in the baseline. In Table 8, the associated coefficients were small but some of them were statistically significant. However, the LATE estimate for the model excluding wage income (at $t-1$) of the wife and husband (variation c) is almost identical to the baseline. The same is true when we exclude hospitalization (at $t-1$) of the wife and husband (variation d). This shows that our LATE is robust to excluding some of the control variables that may be deemed important on a priori grounds. Those, at least, do not cause (additional) omitted variable bias.

B.3.4 Age polynomial, age group dummies

Our statements concerning the probability to retire are to be understood conditional on wife's and husband's age. Clearly, individual age is an important determinant of one's own retirement status. This may make our result sensitive to the specification of the age function used in our model. We control for wife's and husband's age fixed effects in the baseline. As a robustness check, we instead use a third-degree polynomial age function.[20] The LATE estimate for the alternative model is similar to the LATE estimate for the baseline model (variation e).

The difference-in-difference estimates in Table 5 are based on a somewhat less flexible specification than the one we use in our IV baseline model. The main difference was that we used a simple interaction between the treatment group dummy and the year 2005 as an instrument rather than a further interaction with the husband's age dummies. We now test if this latter interaction is necessary. Using a simple single instrument shows that the LATE is similar to (but slightly higher than) the baseline model (variation f). The same holds if we estimate the second stage corresponding to the first stage as shown in Table 6 (variation g). In Table 6 we employed a slightly more flexible functional form than in Table 5, as we used the interactions of dummies for the husband's age categories 55–57 and 58–60 and the husband being employed as a civil servant and the year 2005 as the instruments.

19. The industry dummies are based on the 1993 version of the NACE classification: Agriculture, forestry and fishery, manufacturing, construction, retail, health care, catering and retail, transportation and communication, asset management, commercial services, education, temporary work (reference category: other services).

20. Nonlinear husband's/wife's age effects enter the model through the terms (*husband's age* − 54), (*husband's age* − 54)^2, (*husband's age* − 54)^3, (*wife's age* − 31), (*wife's age* − 31)^2, and (*wife's age* − 31)^3.

TABLE B2
Placebo Test (IV Estimates)

	First Stage Retirement Status of the Husband			Second Stage Retirement Status of the Wife		
	Coefficient	**Std. Err.**	***p* Value**	**Coefficient**	**Std. Err.**	***p* Value**
Predicted retirement status husband				0.1071	0.1978	0.588
Placebo instruments						
Year 2005 × Husband no civil servant × Husband's age 55	−0.0103	0.0033	0.002			
Year 2005 × Husband no civil servant × Husband's age 56	−0.0023	0.0039	0.566			
Year 2005 × Husband no civil servant × Husband's age 57	0.0053	0.0048	0.277			
Year 2005 × Husband no civil servant × Husband's age 58	0.0054	0.0057	0.345			
Year 2005 × Husband no civil servant × Husband's age 59	−0.0040	0.0064	0.529			
Year 2005 × Husband no civil servant × Husband's age 60	−0.0067	0.0127	0.597			
F statistic instruments	14.68					
N	95,141					

Notes: See Table 8 for estimator, and baseline sample. In deviation from Table 8, the specification uses interactions of private sector workers with age and year 2005 as instruments, and includes the interaction of civil servants and the year 2005 as control variable. Other regressors as in Table 8.

TABLE B3
Robustness Checks on Functional Form Specification (IV Estimates)

Variation	**LATE**	**Std. Err.**	***p* Value**	***F* stat.**
Baseline (Table 8)	0.1042	0.0372	0.005	40.16
a. Wife's retirement within 24 months	0.0941	0.0502	0.061	40.16
b. Incl. industry dummies for wives and husbands	0.1032	0.0370	0.005	40.60
c. Excl. wife's, husband's wage income [$t-1$]	0.1051	0.0373	0.005	40.00
d. Excl. wife's, husband's hospitalization [$t-1$]	0.1043	0.0373	0.005	40.11
e. Third degree age polynomial for both spouses	0.0950	0.0361	0.008	41.35
f. Single instrument: husband is civil servant in 2005	0.1212	0.0399	0.002	202.25
g. Two instruments: husband is civil servant in 2005 and husband is aged 58–60 and civil servant in 2005	0.1092	0.0383	0.004	115.58
h. Third degree calendar year polynomial	0.1059	0.0373	0.005	39.98
i. Random effects	0.0911	0.0370	0.014	237.00†

Notes: See Table 8 for estimator, specification, and baseline sample. Variations alter, one by one, the functional form of the regression equation (regressor set), without changing the sample selection (variations a–h), or control for random effects (RE) (variation i). The first stages of variations f and g correspond to the models presented in Tables 5 and 6. †For the RE (variation i), we display a Wald test statistic for joint significance (χ^2 distributed).

B.3.5 Year polynomial

The way we control for variation in retirement rates over years may affect our estimates as well. We control for year fixed effects in our baseline model. Using a third degree polynomial in year instead does not affect the LATE (variation h), however.[21]

B.3.6 Random effects

We also estimate the IV model allowing for individual random effects (RE) to verify whether individual time-constant unobserved heterogeneity affects our result. We need to make the strong assumption that the RE terms are orthogonal to the included regressors in both equations of our IV model. Consistent RE estimates may be characterized by higher efficiency (lower standard errors), Table B3 (variation i) shows that the LATE estimated using the RE model drops slightly compared to the baseline. The significance drops too, however, possibly pointing at violation of the RE assumptions.

21. The polynomial terms are included as ($year-2000$), $(year-2000)^2$, and $(year-2000)^3$.

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